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Turnover Among Young Flemish Managers

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I. INTRODUCTION

Each year, many young people graduate from college and enter the job market with a lot of ambitions, high expectations, and an eagerness to succeed. For an organization, they form an important pool of potential employees. However, turnover during the initial employment period is known to be high. de Pascale and Lange (1971) report that over 50 percent of all American MBA's leave their first employer within five years. Similarly, Sonnenfield, Peiperl and Kotter (1988) find that a majority of Harvard MBA's work for more than two organizations in their first ten years after graduation. In Flanders, 33 percent leave their first employer, while 20 percent has already worked for three or four organizations, within two years after graduation (De Personeelsgids (1991)).

This high turnover during the initial employment period bears implications for both parties involved. From an individual perspective, the early years on one's first job can be difficult. Young graduates often have unrealistic job expectations, may experience ambiguity about the criteria by which their performance will be evaluated, or may be unaware of the organization's political aspects. Because of these and other reasons, the initial employment period is often characterized by frustration and dissatisfaction leading to a high turnover rate. Moreover, events in the early stages of one's career have been found to be important determinants of socio-economic attainment in subsequent stages (see e.g. Shavit, Matras and Featherman (1990)), which amplifies the implications of the phenomenon. From an organizational

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perspective, high turnover in the initial employment period can become very costly. Besides the obvious costs of recruitment and training, organizations may be confronted with a low morale of the other employees, interruption of the work flow, and damage to their reputation. For them, it is important to monitor turnover rates and to take corrective action when the implications of turnover become excessive.

The development of strategies for dealing with turnover rates depends on the level of understanding of the issue under investigation. However, this understanding may be biased by the methodology used in most research and by an excessive reliance on U.S. data. As to the methodology, we use the flexible hazard-rate specification developed in Vanhuele, Dekimpe, Sharma and Morrison (1995). Unlike regression models (where the dependent variable is the observed length of stay with the company) or logit/probit models (where one defines a zero-one variable to indicate who has left the company by the end of the observation period), the hazard function is our construct of main interest. This function measures the conditional probability of leaving the company in a given time period t, given that one has not yet done so by the end of period t-1. As explained in more detail below, the hazard formulation used in this paper has the following appealing features: (1) it allows for right-censored data (i.e. some people may still be on their first job at the end of the observation period, and their true duration will only be known in the future); (2) it incorporates industry-specific, firm-specific and individual-level characteristics, which have been suggested in previous research as potential moderators of one's propensity to quit; (3) it measures the impact of tenure in a non-parametric way, and thereby allows for non-monotonic patterns in the relationship between an individual's quitting probability and his/her length of stay with the organization, and (4) it corrects for the fact that our model does not yet incorporate all potentially relevant factors such as the amount of extra-legal benefits. While previous studies have already examined predictors and reasons for leaving an organization (e.g. Terborg and Lee (1984); Cotton and Tuttle (1986); Lewis and Park (1989); Petersen and Spilerman (1990); Stroh, Brett and Reilly (1993)), their methodology did not incorporate all aforementioned properties. Still, each property is important to make reliable and managerially useful inferences from the data at hand. Second, most research on turnover has been conducted in the U.S. Recently, the need to assess the appropriateness of transferring theories developed in the U.S. to other countries has been emphasized. Given the growing awareness of the cultural relativity of modern organization theories, testing prevailing theories on non-U.S. data is increasingly encouraged (see e.g. Berry (1989); Boyacigiller and Adler (1991); Hostede (1983)).

Thus far, little empirical research has been conducted on the job mobility of university graduates in Flanders. A rich data set has been made available to us by *Krantengroep De Standaard*, on the basis of which we assess the following issues:

- How long do Flemish graduates, on average, stay with their first employer, and when are they most likely to quit?
- Are female graduates more or less likely to switch jobs in an early stage of their career?
- What career motives are valued most by early movers, extrinsic (such as salary) or intrinsic (such as autonomy)?
- Do small firms have a competitive disadvantage compared to larger firms because of a lack of promotion possibilities, or does their more entrepreneurial character encourage commitment?

The remainder of the article is structured as follows. In Section II, we formulate our hypotheses with respect to the impact of tenure, education, gender, geographical preference, career values, salary and size of the organization on a graduate's quitting probability. Section III describes our data, and Section IV discusses briefly the adopted modeling approach. Empirical results are presented in Section V. Section VI summarizes our main conclusions.

II. ANTECEDENTS OF TURNOVER[1]

A. *Tenure*

The relationship between an individual's length of stay with the organization (his/her tenure) and the probability of quitting has been the subject of much research interest. Terborg and Lee (1984) found a negative relationship between the time in an employee's present job and the turnover rate, implying that the longer someone has been with the company, the lower his/her propensity to quit. Based on a meta analysis of previous studies, Cotton and Tuttle (1986) also reported a highly significant negative relationship between tenure and turnover. An underlying reason for this relationship is that tenure is a proxy for the amount of firm-specific training or specialization which presumably is of less value to other firms (Blossfeld and Hamerle (1990); Mincer (1974); Mobley (1982)). Also, the longer an individual has been with a particular organization, the more he/she may become emotionally attached to it and perceive a congruence between his/her goals and those of the organization (Huselid and Day (1991)).

Alternatively, Blossfeld and Hamerle (1990) report a positive duration dependence at the individual level after controlling for unobserved heterogeneity (see Section IV for a discussion on this issue). In other words, they find that employees become more likely to quit as they have spent more time in their current job.

Third, one could hypothesize a non-monotonic relationship between tenure and turnover. Indeed, young graduates may accept the first job offer they get and use that job to look in a more relaxed way (since they are no longer unemployed) for their "ideal job". Because of this "job shopping", one may see an increasing quitting probability during the first months/years after graduation. Also, new entrants into the labor market face significant occupation- and job-specific uncertainty (McCall (1990)), and may start their first job with unrealistic expectations. Again, this may cause the quitting probability to increase during the first months or years. However, if they "stick around" for some time, the congruence with the organization may eventually increase, their firm-specific training may start to accumulate, and their quitting probability may, past a critical duration, start to decrease. Such a non-monotonic pattern was found by McCall (1990), who concluded that models which only allow for monotonically increasing or decreasing relationships between turnover and tenure may not be appropriate.

Our first hypothesis is therefore that a graduate's quitting probability will be affected by his/her length of stay on the job. However, the nature of this relationship is not clear a priori. As such, we will model the relationship non-parametrically, which offers us enough flexibility to also capture non-monotonic and even multi-modal patterns.

Hypothesis 1: The probability of quitting the first job will be related to the elapsed time in that job.

B. *Education*

According to human capital theory, workers make rational choices regarding investments in their own human capital (Becker (1975)), including education, interruptions in labor force participation and company changes. Other things being equal, managers with greater human capital may be more likely to exit their organizations than managers with less human capital. The primary indicator of human capital, education, is generally found to be positively correlated with turnover (Cotton and Tuttle (1986); Terborg and Lee (1984)). Organizations, on the one hand, may use education as an indicator of ability and consider managers with higher levels of education more easily reemployable than those with less education. Managers, on the other hand, may change companies as a way to increase their human capital by obtaining breath of experience. For example, career handbooks (e.g. Gerberg (1980); Hirsch (1987)) advise people to change organizations if they desire substantial salary changes. This advice seems predicated on the belief that only in the external labor market will one obtain one's full value. Skills are presumably insufficiently rewarded by the employing organization in the absence of competitive market forces.

Hypothesis 2: Education will be negatively related to duration in the first job.

C. *Gender*

Female employees are often believed to have higher turnover rates than male employees (Cotton and Tuttle (1986)). Schwartz (1989), for example, found that top performing females have turnover rates that are nearly two and a half times that of their male counterparts. Similarly, Lewis and Park (1989) report that the percentage of quitting female administrators and professionals in federal service is higher than that of males. On the other hand, Stumpf and Dawley (1981) found males to be more likely to leave than females in female-dominated nonprofessional occupations, and Flanders and Anderson (1973) and Hennig and Jardim (1977) found that female managers were more inclined to have single-employer careers.

A potential explanation for these conflicting findings lies in the sample-specific causes for quitting. Petersen and Spilerman (1990) found that women quit their company less often than men for career reasons, but have a much higher quitting rate for personal reasons. Thus, if women often appear to have a lower job attachment than men, it is not that they move more from one employer to another, but because they either quit the job-market to take care of the children or because they follow their husband after he has changed jobs. Since we do not have firm priors on the importance of these family obligations in our data set (i.e. shortly after graduation), we do not formulate a directional hypothesis with respect to the impact of gender.

Hypothesis 3: Gender will be related to duration in the first job.

D. *Geographical Preference*

Job mobility may also be a result of self-selection. The self-selection argument suggests that individuals, by choice, choose for a specific career path. Brett, Stroh and Reilly (1992) found that self-selection variables such as willingness to move, attitudes toward moving, number of times they turned down a transfer, and number of times their name was removed from consideration for a transfer accounted for a significant amount of the variance in both salary progression and geographic mobility. Likewise, employees who were mobile before, continue to be more mobile (Lindeboom and Theeuwes (1991)). This study examines preference of job location as an indicator of self-selection, expecting that individuals with no specific geographical preference will leave their first employer sooner.

Hypothesis 4: Having a specific geographical preference will be positively related to one's duration in the first job.

E. *Career Values*

Workers enter jobs with different work and career values, which increases the heterogeneity of the work force (London and Stumpf (1986); Morris and Villemez (1992)). Since career-related variables describe the employee population, they may also be important factors in turnover and job-mobility research. Recent research found that the major factor predicting turnover for both male and female managers was their intent to leave, which was in turn predicted by attitudes about the job, company and career (Stroh, Brett and Reilly (1993)). Managers who were satisfied with their jobs intended to stay longer, as did managers who were loyal to the organization. On the other hand, managers who were self loyal and willing to change companies for career advancement, were particularly likely to leave.

This study examines the relation between duration in first job and four different career values: job initiative, job compensation, career advancement and organizational status[2]. These career values represent the young graduates' attitudes towards work and career when choosing their first job. While job initiative refers primarily to an intrinsic motivational value to accept a job, job compensation, career advancement and status refer more to different components of extrinsic motivational values. Since the influence of these career values on turnover is examined in an exploratory way, no directional hypotheses are formulated.

Hypothesis 5: Individual career values will affect the duration in the first job.

F. *Salary*

Salary is one of the most salient job characteristics related to turnover. According to neoclassical labor market theory, an organization which offers high net advantages will have a more stable labor force than organizations which offer low net advantages as advantage-maximizing employees seek employment elsewhere. Put differently, since salary level is one of the most measurable components of net advantage, individuals with high salaries are expected to have less incentive to move (McCall (1990)). For example, Curran (1981) found that employees in industries with relatively low wages have a greater propensity to quit their job than those in industries with higher wages. The meta-analysis of Cotton and Tuttle (1986) confirmed the conclusion that actually-paid salary is a stable, reliable predictor of turnover.

Hypothesis 6: Salary will be positively related to the duration in the first job.

G. *Type of Organization*

The major indicator of type of organization we consider is their size as measured by their number of employees. Previous research indicates that the size of an organization is related to its turnover rate. However, the direction of this relation is not always clear (Mobley

(1982); Terborg and Lee (1984)). Large organizations may have lower turnover rates because of greater internal mobility opportunities, more sophisticated human resource management processes, more competitive compensation systems and more activities devoted to the management of turnover. Since the internal labor market is more efficient, people will stay with the same company. Similarly, if the rewards within the organization are already quite high, it becomes less likely that the opportunities outside the company are more attractive (Petersen and Spilerman (1990)). However, one could also argue that large organizations may experience a higher turnover rate because of communication problems, poor group cohesion, impersonalization and bureaucratization. So, small organizations may have a lower turnover rate since they may be regarded as having a more congenial working atmosphere and more opportunities for initiative. Finally, there is some suggestion that size might be curvilinearly related to turnover with small and large organizations having the least turnover (Bluedorn (1982)).

The size of an organization may be a reasonable descriptor for private companies. However, several graduates in our sample had jobs in the governmental and educational sector, for which size is a less meaningful descriptor. For those categories, a separate dummy variable was introduced, but no a priori hypothesis on the sign of their coefficients (relative to the turnover rate in the largest private-company category) was formulated.

Hypothesis 7: The type (size) of company will be related to the duration in the first job.

III. DATA

A. *Sample*

The sample consists of 704 young Belgian managers. They graduated in 1986 and 1987 from a Flemish university or industrial college. All of them were employed for at least 24 months, although not necessarily with the same employer. The duration of interest in our analysis is their length of stay with their first employer. For those who had not yet switched jobs, the duration until the end of the observation period (June 1990) was considered. As will be discussed in section IV, our modeling approach will explicitly account for this non-switching.

Sixty six percent of them were male, 34 percent female. Their education degree was in business (22%), engineering (18% industrial engineering and 10% civil engineering), literature (16%), law (12%), human sciences (11%), and exact sciences (11%). The jobs covered mostly administrative (19%), educational (18 %), research related (15%), commercial (14%) and technical (8%) functions. Other jobs were in information technology, production, personnel, and marketing. The industry in which they were employed was mainly education (16%), business services (12%), and banking and insurance (11%).

B. *Measurement*

Table 1 presents the different categories of each explanatory variable, and gives the number of observations in each category. Except for career values, all variables were coded as dummy variables.

Career values were measured by asking the respondents to rate the importance of 14 criteria in choosing their first job on a 5-point scale ranging from not at all important (=1) to very important (=5). Con-

TABLE 1
The categories of explanatory variables

Explanatory Variable	Categories	Frequencies (n = 739)	
Education			
Type of Degree	Industrial College *	177	
	University	562	
Extra Degree	No *	422	
	Yes	317	
Gender	Male *	491	
	Female	248	
Geographical Preference	Preference *	251	
	No Preference	488	
Gross Salary	< BF 45,000 *	208	
	BF 46,000 - BF 60,000	387	
	> BF 60,000	179	
Type of Organization	> 250 Employees and International *	249	
	51 - 250 Employees	89	
	< 51 Employees	102	
	Government	105	
	Education	151	
	Other	43	
		mean	standard deviation
Career Values	Job Initiative	28.42	6.11
	Career Advancement	15.52	3.81
	Job Compensation	10.92	2.61
	Job Status	7.32	2.29

* defines the base case in the hazard rate analysis

firmatory factor analysis using Lisrel was used to examine the unidimensionality of the four career values identified (using exploratory factor analysis) in the original study of *Krantengroep De Standaard* (De Personeelsgids (1991)): job initiative, career advancement, job compensation, and status. After respecification, the measurement model included 12 items and provided an acceptable fit ($C^2 = 344.9$, $p < .001$; GFI= .94). The items and their factor loadings in the confirmatory factor analysis are presented in Table 2. Subjects' scores on each career value were then computed as the weighted linear composite of the item scores (Steenkamp and van Trijp (1991)).

TABLE 2
The measurement model of four career values

	Factor Loadings	Alpha
Job Initiative		
Freedom to take initiative	.86	.82
Having responsibility	.77	
An entrepreneurial job	.68	
Job with autonomy	.63	
Career Advancement		
Opportunities for career advancement	.82	.67
Opportunities for training	.61	
Opportunities for job rotation	.50	
Job Compensation		
Good salary	.55	.66
Fringe benefits	.67	
Compensation on the basis of results	.67	
Job Status		
Job with hiçgh personal status	.70	.68
Job in a high status organization	.68	

IV. THE MODELING APPROACH

To test the hypotheses introduced in Section II, one may be inclined to use a regression model with the number of months in the first job as dependent variable, and the respondents' gender, education level,... as explanatory variables. However, the distinct nature of our duration data makes this approach less appropriate. First, 43 percent of the respondents were still in their first job at the end of the observation period. As a consequence, their true length of stay before switching will only be known in the future. Such observations are called right censored and cause severe difficulties in a regression context. One way of dealing with censored observations is to ignore them, i.e. to limit the sample to those respondents who have already left their first job and for which we know the actual length of stay. However, this would result in a significant loss of information and would cause the parameter estimates to be biased. Alternatively, one could use the observed duration (i.e. until the end of the observation period) as a proxy for their unknown true duration, but this would again lead to biased and inconsistent parameter estimates (see e.g. Lancaster (1990)). Second, if one wants to assess the impact of tenure on turnover, the length of stay should be used as an explanatory rather than as dependent variable. Put differently, the nature of the time dependence cannot be assessed with traditional regressions. Third, even when many potentially important moderating variables are included into the model, there may still be factors which are omitted from the model. For example, the amount of fringe benefits offered by the company, the distance between the company and the respondent's home, or the relationship with his/her co-workers may all have an impact on a young professional's propensity to quit, but no data on these factors were available in the data set. The omission of these potentially important factors may affect the included parameter estimates, but it is difficult to correct for an omitted-variable bias in a regression context[3].

One way to deal with censored observations is to use a logit model, where the dependent variable can take on two values, 0 for those who have not yet quit their first job (the "stayers") and 1 for those who have left (the "switchers"). From this logit estimation, one can asses what factors affect a respondent's probability of belonging to either category. However, this again results in a significant loss of information, since no distinction is made between someone who changed jobs after 3 months and someone who switched after 3 years; both will be coded as "switchers". As will be indicated in our empirical results section, this loss of information may have substantial implications when testing our hypotheses. Also, it is still difficult to correct for unobserved heterogeneity and to assess the nature of the time dependence (see Gupta (1991) or Helsen and Schmittlein (1993) for a detailed discussion).

The proposed framework will adequately deal with censored observations, in that it will use all relevant information on those who have not yet left their first job. It will also allow us to quantify in a very flexible way the nature of the time dependence, and will provide a correction for unobserved heterogeneity. In what follows, we only present the spirit of the model, and refer to Vanhuele, Dekimpe, Sharma and Morrison (1995) for a detailed technical discussion.

A. *The Base Model*

A central concept in our modeling approach is a respondent's hazard, i.e. the conditional probability of changing jobs in period t given that one has not done so earlier. The form of the hazard function will reveal the impact of tenure (also called the nature of the time dependence) and will indicate, for example, whether someone who has been on the job for three years is more, less or equally likely to quit in the coming period than someone who has been on the job for only two years.

The most simple starting point is to assume that there is no time dependence, i.e. irrespective of how long a young professional has been on the job, he/she is equally likely to quit in the coming period. Technically, this is reflected in the assumption that the manager's duration has an exponential distribution. Indeed, for the exponential distribution the probability density function $f(t)$, cumulative distribution function $F(t)$, survival function $S(t)$ and hazard function $h(t)$ are given by[4]

$$f(t;\lambda) = \lambda\ e^{-\lambda t}, \tag{1a}$$

$$F(t;\ \lambda) = 1 - e^{-\lambda t}, \tag{1b}$$

$$S(t;\ \lambda) = 1 - F(t;\ \lambda) = e^{-\lambda t}, \tag{1c}$$

$$h(t;\ \lambda) = \frac{f(t;\ \lambda)}{1 - F(t;\ \lambda)} = \lambda\ . \tag{1d}$$

As shown in Equation (1d), the conditional probability of quitting (the hazard) is a constant, and therefore does not depend on the time the respondent has been with the company.

To obtain an estimate of the parameter λ, we derive the (log) likelihood function, which is subsequently maximized with respect to λ.

The contribution of a respondent to the likelihood function is different for completed and censored observations. The likelihood of seeing a completed duration of length t_i is given by $f(t_i)$. However, to account for the discrete nature of the data (we know in what month he/she has quit the job, but do not know the exact day within that month), we replace $f(t_i)$ by the difference between the survival function at the beginning and end of the month. As such, the contribution becomes $S(t_i\text{-}1)\text{-}S(t_i)$, where the survivor function $S(t_i)=1\text{-}F(t_i)$ gives the probability that manager i will stay at least t_i periods. Consider, for example, an exponential model where λ equals 0.2. Using equation (1c), $S(1)$ can be computed as $e^{-0.2}$ (= 0.82), and gives the probability that the manager stays *at least* one period. Similarly, $S(2)$ (= $e^{-2*0.2}$ = 0.67) gives the probability of staying *at least* two periods with the company. The difference between these two probabilities (= 0.15) gives the (unconditional) probability that the manager leaves *during* the second period. For a censored observation, the contribution to the likelihood function is given by $S(t_i\text{-}1)$. We therefore assume the censoring to occur at the beginning of the time interval $[t_i\text{-}1,t_i)$. Clearly, some assumption of this kind is needed given the discrete nature of our data. Note that this likelihood contribution reflects all relevant information we have on this person, i.e. that he/she had not yet left after the observed number of periods.

If we introduce a censoring dummy d_i which takes the value of one in case of a censored observation and zero otherwise, we can write the likelihood contribution of manager i as:

$$L_i(t_i;\,\lambda) = [S(t_i - 1) - S(t_i)]^{1-d_i}\,[S(t_i - 1)]^{d_i}\,. \tag{2}$$

If the managers' durations are exponentially distributed, and if we assume that they all have the same mean quitting rate λ, the log-likelihood function for a sample of N managers becomes:

$$LL = \sum_{i=1}^{N} (1-d_i)\ \ln[e^{-\lambda(t_i-1)} - e^{-\lambda_i t_i}] - d_i\,\lambda\,(t_i - 1)\,. \tag{3}$$

B. *Incorporating Observable Characteristics*

Thus far, we have made the unrealistic assumption that all young professionals have the same (conditional) quitting probability. In this section, we relax that assumption, and allow graduates with different characteristics to have different quitting rates. This will allow us to test most[5] of the hypotheses discussed in Section II. Following Vanhuele et al. (1995), we write the quitting rate of the *i*th manager as[6]

$$\lambda_i = \lambda_0 \, e^{b X_i},$$

X_i is a vector of explanatory variables, b is a vector of coefficients, and λ_0 is the quitting rate of the base group which comprises those graduates for which all covariates are zero (see Table 1 for a description of the base category). How should one interpret the b-coefficients? A positive coefficient implies that an increase in the value of the covariate increases the conditional probability of quitting in the next period relative to the base group (and thus reduces the expected length of stay). Moreover, after some algebra one can show that when the j-th covariate changes by one unit, the hazard function changes by $[\exp(b_j)-1]$ *100 percent. To derive the likelihood function (which is subsequently maximized with respect to λ_0 and b), an expression for the survivor function associated with the hazard rate in equation (4) is needed. This expression is derived in Vanhuele et al. (1995), and we refer the interested reader to the original article for mathematical details.

C. *Modeling the Impact of Tenure*

Even though in Section B we already allowed managers with different observable characteristics to have a different quitting propensity, we still assumed that the hazard did not change over time. To capture the effect of tenure, we add a set of time-varying dummy variables to equation (4), which becomes (Vanhuele et al. 1995)

$$\lambda_i(t) = \lambda_0 \, e^{b X_i} \, e^{c D_i(t)} .$$

Theoretically, a separate dummy variable could be used for each month the manager has been in his/her first job. For example, the time-varying dummy associated with month 3 is always zero except during the third month when it takes the value of one, and therefore becomes (0 0 1 0 0... 0). Similarly, the variable associated with month 4 becomes (0 0 0 1 0 0... 0). No dummy variable is included for the first month, since the joint estimation of λ_0 and c_1 would result in identification problems. λ_0 should therefore be interpreted as the quitting rate of the base group in the first month. Positive (negative) c-coefficients for the other months indicate a higher (lower) quitting probability as compared to the first period. In our empirical application, we will not allow for a different c parameter in every month, since this would consume too many degrees of freedom. Instead, we will allow for a different c-parameter after 3, 6, 12, 18, 24, 30 and 36 months.

Intuitively, this approach can be interpreted as a piecewise approximation to an underlying, possibly very complex, hazard function describing the nature of the time dependence. Note that we do not make any distributional assumptions on the form of the time dependence, as opposed to the exponential model which implies no time dependence or the Weibull model used in Lindeboom and Theeuwes (1991) where only a monotonically increasing or decreasing time dependence can be captured. We therefore refer to our approach as a non-parametric estimation of the baseline hazard (which is the hazard of the base group). From a substantive point of view, its flexibility allows us to quantify the impact of tenure without imposing stringent and often hard-to-justify a priori assumptions. From a statistical point of view, this method ensures the consistency of the parameter estimates even when the true underlying continuous hazard is unknown (Meyer, 1986, 1990).

We again refer to Vanhuele et al. (1995) for a derivation of the survivor function (which is then substituted into equation 2) associated with the hazard function in equation 5.

D. *Accounting for Unobserved Heterogeneity*

Some factors that affect a manager's propensity to quit may not be available in the data set at hand or may be hard to quantify. Not accounting for these omitted factors may cause (1) a spurious negative duration dependence (i.e. the estimated c-parameters will be too small, which may lead to misleading inferences about the impact of tenure), and (2) will also cause the coefficients of the included covariates (which are used to test our hypotheses) to be biased and inconsistent (see Vanhuele et al. (1995) or Dekimpe and Degraeve (1995) for an elaborate discussion on this issue).

In a nutshell, we let λ_0 be distributed according to a certain mixing distribution. As such, even though all managers in the base group have then the same observable characteristics, they are still allowed to have a different quitting propensity. To derive the mean baseline hazard, we subsequently use the mean of the mixing distribution (where for mathematical convenience we use the gamma distribution). The interpretation of the c and b-coefficients is not affected by this operation; they still reflect how the hazard shifts relative to, respectively, the first month and the base group managers.

In the discussion of the empirical results, we will present parameter estimates for a model with and without a correction for unobserved heterogeneity. As such, we will be able to clearly illustrate the downward bias on the estimated time dependence when not accounting for this phenomenon.

TABLE 3
Parameter estimates

	Logit	Without Heterogeneity (1)	With Heterogeneity (2)
Tenure			
C2 (4 - 6 months)		0.618***	0.645***
C3 (7 - 12 months)		0.642***	0.709***
C4 (13 - 18 months)		0.804***	0.929***
C5 (19 - 24 months)		0.812***	0.990***
C6 (25 - 30 months)		0.420*	0.641*
C7 (31 - 36 months)		0.436	0.680
C8 (37 - 54 months)		-0.744	-0.448
Education			
Type of Degree	0.087	0.050	0.513
Extra Degree	-0.025	0.064	0.075
Gender	0.027	0.036	0.044
Geographical Preference	0.479***	0.299***	0.329***
Career Values			
Job Initiative	-0.063***	-0.038***	-0.043***
Career Advancement	-0.037	-0.026*	-0.026
Job Compensation	-0.044	-0.013	-0.014
Job Status	0.056	0.030	0.033
Gross Salary			
BF 46,000 - 60,000	-0.447**	-0.167	-0.187
BF > 60,000	-0.947***	-0.567***	-0.615***
Type of Organization			
51 - 250 Employees	0.882***	0.588***	0.647***
< 51 Employees	0.391	0.176	0.197
Government	0.512*	0.387**	0.436**
Education	0.803***	0.448***	0.509***
Other	-0.072	-0.092	-0.091

(1) $\lambda_0 = 0.044$

(2) the mean of the mixing distribution is 0.047

*** significant at the 1% level

** significant at the 5% level

* significant at the 10% level

V. RESULTS AND DISCUSSION

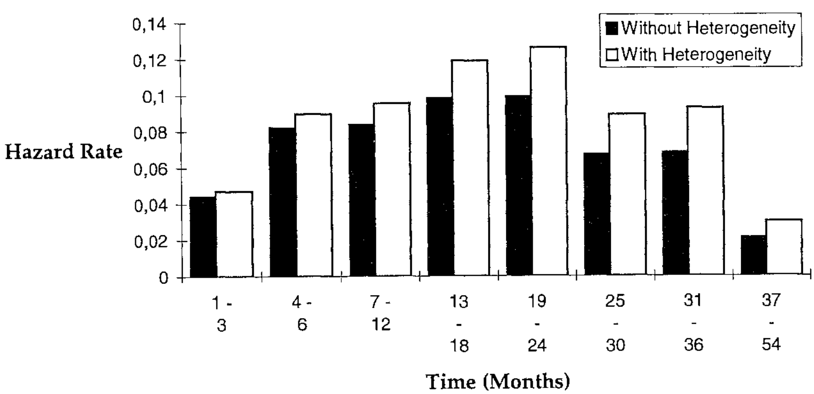
Maximum-likelihood estimates for the model with and without heterogeneity are given in Table 3. For comparison purposes, we also present parameter estimates for the more familiar logit model.

A. *Tenure*

To assess the impact of tenure, we plot in Figure 1 the baseline ha-

FIGURE 1

The impact of tenure on the hazard rate of young Flemish managers



zard. Remember that the baseline hazard gives the conditional quitting probability for managers in the base group (i.e. male graduates from an industrial college, without an extra degree, with a specific geographic preference, earning less than BF 45,000 as starting salary, working in an international firm with more than 250 employees, and with a score of zero on the career values). For the other categories, the hazard is proportional to this graph, and is obtained by multiplying the baseline hazard with exp[bX], where X is their vector of covariate values. Two issues need to be emphasized with respect to Figure 1. From a substantive point of view, we see that the nature of the time dependence is non-monotonic and increases in the first two years, after which a decrease in the quitting probability is observed. As such, our findings confirm the results of McCall (1990) who also found a non-monotonic impact of tenure. From a methodological point of view, Figure 1 clearly illustrates the downward bias on the estimated time dependence when not accounting for unobserved heterogeneity. Put differently, if one does not explicitly correct for the fact that not all relevant factors can be included into the model, one will seriously underestimate the employees' quitting probability, which may in turn have a negative impact on the reliability of aggregate turn-over plans. Also, the pattern of the time dependence in Figure 2 may be hard to capture with a parametric distributional form, which illustrates the desirability of adopting the extremely flexible, non-parametric specification.

B. *Education*

In contrast to hypothesis 2, the results showed no relation between first job mobility and education. Flemish managers with a university degree or with an additional education did not leave their first employer sooner than those with a degree from an industrial college, or no additional education. Two phenomena may explain this disconfirmation of our prior hypothesis. First, the hypothesis was based on human-capital theory assuming an efficient external labor market. Our findings may be due to the fact that the internal labor market is an important factor in Belgium. Leaving one's job may cause the job-specific expertise to become less useful, and correspond to a drop in their built-up human capital. Second, having a degree from an industrial college as opposed to a university degree, or the acquisition of a second degree, may not induce much difference in initial human capital. Put differently, the entire sample is well educated, and the resulting variability along the human-capital dimension may not be large enough to find significant differences for the education variable.

C. *Gender*

The results also showed no relationship between gender and first job mobility (hypothesis 3). It is possible that gender differences with respect to turnover are not encountered in the early stages of one's career. de Pasquale and Lange (1971) also found that men and women were equally likely to move from one organization to another in the early years of their careers. As suggested in Section II, one explanation may be that the female graduates in our sample had the same (lack of) family obligations as their male colleagues. An alternative explanation may be that in the beginning of their career men and women have similar work and career related experiences. Indeed, Miller and Wheeler (1992) and Stroh, Brett and Reilly (1993) found that gender differences in turnover disappear after controlling for these experiences. In a recent study in The Netherlands, Lindeboom and Theeuwes (1991) found that females are more mobile than males, but that the observed differences were partly due to males and females having differing job characteristics (e.g. males work on average 38 hours, whereas females work an average of 25 hours a week). In our sample of young professionals, we do not expect such dramatic differences to exist, which may explain why no significant differences in turnover were found among the two groups.

D. *Geographical Preference*

Flemish managers with no preference for a specific job location were found to leave their first employer sooner than the base group. All else equal, the conditional quitting probability of a manager who has no specific preference where to work is 39 percent higher than of a manager who has a specific preference. This supports our hypothesis that geographical preference is an important self-selection variable in understanding the job mobility of Flemish managers.

E. *Career Values*

Our results show that Flemish managers who value job initiative more are less likely to leave their first job. There were no differences found with respect to the stated importance of job compensation, career advancement and job status. Young managers who chose their first job based on criteria like freedom to take initiative, opportunity to take responsibility, and autonomy, had a lower quitting probability (i.e. 4% = [exp (-,043)-1)*100]) than those who valued job initiative lower[7]. Other evaluation criteria like a good salary, a job with career advancement, and with high personal status had no impact on the duration in the first job. These results are consistent with Kraut's study (1975) in which attitude towards work was found to be significantly correlated with turnover but not attitude towards company, team work, advancement, pay and workload. The distinction between these two types of criteria seem to refer to intrinsic and extrinsic motivation. This explorative finding suggests that managers who are intrinsically motivated are more likely to stay in their first job.

F. *Salary*

We find that Flemish managers with a starting salary of more than BF 60,000 are less likely to leave the organization than those with a salary of less than BF 45,000. The former group's conditional quitting probability is 46 percent [exp (-.615)-1)*100] lower than that of the base group. This finding supports the general conclusion that pay is a stable and reliable predictor of turnover. It is worth noting that even though the respondents' *initial (and stated)* interest in monetary compensation turned out to be insignificant (cf. supra), that their *actual* salary does affect their expected length of stay. A further investigation of this discrepancy is certainly an interesting area for future research.

G. *Type of Organization*

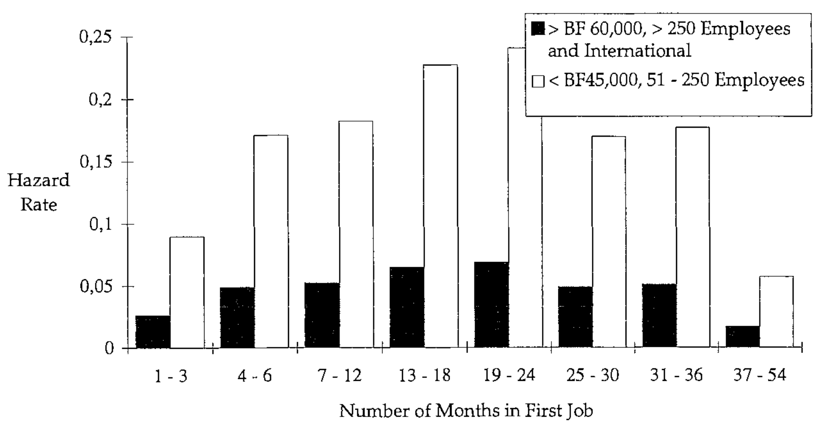
Flemish managers working in medium large organizations (51›250 employees) have a 90 [exp(.647)-1)*100]percent higher quitting probability than those working in a large, international organization. No differences in first job mobility were found between managers working in, respectively, the latter type of organization and small organizations with less than 51 employees[8]. This finding provides support for a curvilinear relationship between turnover and size of the organization. Turnover is the highest in medium-sized organizations which may not have the internal mobility opportunities of a large organization nor the congenial working atmosphere and opportunities for initiative of a small organization.

This study also showed that Flemish managers who started working in a governmental and educational organization leave this organization sooner than those in a large private organization. Their conditional quitting probability is respectively 58 percent and 66 percent higher than the base group. This finding indicates that first job mobility in Belgium is higher in the public sector than in certain company types in the private sector.

H. *The Quitting Probability of Selected Groups*

To further illustrate how the quitting probability differs quite dramatically among different categories, we plot in Figure 2 the hazard rate for two groups: (1) graduates who receive a starting salary of more than BF 60,000 in a multinational firm, as opposed to (2) graduates who receive less than BF 45,000 in a medium-sized company (55-250 employees). The graphs are obtained by multiplying the top graph in Figure 1 with $\exp(b_j)$ where b_j is the parameter estimate associated with, respectively, the salary or firm-type category. As such, we assume that on all other covariates (e.g. geographic preference, intrinsic mo-

FIGURE 2
The hazard rate of selected groups of young Flemish managers



tivation), both groups have the same values as the base category. Obviously, one could also plot the graphs for other combinations of the different characteristics, and assess the (conditional) quitting probability for different types of employees at different points in their career (i.e. for different levels of tenure).

The results of these analyses can also be presented in terms of the survivor function. As illustrated in Table 4, 92% of the managers in group I stay for at least 3 months in their first job, while almost 25% of the managers in group II have already switched jobs by that time[9].

TABLE 4

The survival probability for selected groups of young flemish managers

	Group I	Group II
S (3 months)	0.927	0.764
S (6 months)	0.801	0.457
S (12 months)	0.589	0.153
S (18 months)	0.400	0.039
...	...	...

Note : Group I : > BF 60,000, > 250 Employees and International
Group II : < BF 45,000, 51 - 250 Employees

VI. CONCLUSION

The probability that young Flemish managers leave their first employer during their initial employment period is influenced by the managers' tenure with the firm, their geographical preference, the importance they attach to job initiative, the actually paid salary, and the size of the organization. In contrast, education, gender, the importance they attach to job compensation, career advancement and job status when choosing their first job, had no effect on the young managers' quitting rate.

Flemish managers show an increasing quitting probability during the first two years of their employment. After these two years, their quitting rate starts to decrease. These findings could not have been obtained with traditional analyses such as regression or logit models. Moreover, Figure 2 clearly illustrated that the time dependence in the managers' quitting propensity would be difficult to operationalize with a parametric distributional form, and would have been seriously under-estimated if no allowance had been made for unobserved heterogeneity. With respect to the other predictors, organizations selecting young graduates may first of all want to focus on the geographic preference of their candidates since this seems to be an important self-selection variable in understanding Flemish managers' job mobility. In addition, young graduates who have chosen their first job based on the intrinsic value of job initiative are more likely to stay longer with their first employer. However, once they have found a job, their actually-paid salary, which is a major extrinsic motivator, also becomes an important determinant of their expected length of stay. Finally, managers working in medium-sized organizations are more likely to leave the organizations than those working in large international or small organizations. They seem to lack the benefits of career opportunities in large organizations, as well as the greater autonomy and congenial atmosphere in many smaller organizations. These findings indicate the importance of assigning young graduates a challenging job, providing them with a career perspective, and creating a stimulating work environment. In taking such actions, organizations are likely to decrease the turnover of their young managers.

NOTES

1. In this paper, turnover is defined as a change of employer. As such, internal promotions are not considered to be a job switch.
2. The operationalization of these concepts is discussed in more detail in Section III.
3. In what follows, we will refer to the fact that not all relevant factors are included as unobserved heterogeneity. This reflects the fact that respondents differ in their quitting propensity, and that not all underlying causes for these differences are observed by the model builder.
4. For notational convenience, we define monthly grouping intervals $[t_{k-1},t_k)$, $k = 1, 2,...,$ $m+1$, $t_0 = 0$, $t_{m+1} = \infty$, and record quitting in duration interval $[t_{k-1},t_k)$ as t_k. It should be emphasized that the variable t does not refer to actual calendar time, but to the number of months in a respondent's first job.
5. A test for the impact of tenure is discussed in Section IV.C.
6. The exponential function is used because it is quite flexible, and because it ensures the non-negativity of the hazard rate.
7. Note that this covariate is not a dummy variable but has a continuous scale. The 4% increase refers to the increase in the conditional quitting probability if the value of the respondent's factor score increases by one unit.
8. Note that this result is in contrast with what is inferred from the parameter estimates in the logit model, which indicates that the information loss in the latter may affect the conclusions of the analysis.
9. The figures in Table 4 are easily derived using the following relationship between the survivor and hazard function: $S_i(t) = \exp[-(\lambda_{i,1}+\lambda_{i,2}+...+\lambda^{i,t})]$ where $\lambda_{i,j}$ is the hazard of manager i in period j (see Vanhuele et al. (1995) for more details).

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